Lifecycle Changes in the Rate of Time Preference:

Testing the Theory of Endogenous Preferences and its Relevance to Adolescent Substance Use

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Abstract

Economic theory indicates that because one's activities to improve health reward one in the future, persons who value the future more highly will be more prone to healthy activity. Without measures of time preference we can neither test this theory nor understand what makes people value future events more highly.

Progress in this area requires a method to infer measures of time preference from the secondary datasets used in public health and economic research. Time preference in its econometric expression is the measurable forfeiting of additional goods in the present to enjoy goods in the future. The rate of time preference varies from 0 for individuals who are indifferent between present and future consumption to infinity for individuals who have place no value on future consumption. When subjects decide to forego higher wages in the present by taking safer jobs that increase their chance of future survival they send signals about their time preference (mixed with signals about risk aversion, other job prospects, family pressures, etc.). These wage-risk tradeoffs offer scholars interested in measuring time preference the convenience of a secondary dataset, but the drawback of needing to control for the confounding and endogenous factors.

This study applies econometric techniques to the National Longitudinal Survey of Youth (NLSY) to derive estimates of the levels of time preference for each labor force participant in each of 15 waves of data from 1979 to 1994. With these estimates I describe the evolution of time preference over the life course. I test the following hypotheses suggested by Becker and Mulligan (Becker and Mulligan 1997)in their theory of endogenous time preferences: 1)Age and Education reduce the rate of time preference; 2)Female gender and Father's Presence in the Home at age 14 reduce the rate of time preference; 3) Religious participation reduces the rate of time preference. Finally I show that subjects with a more immediate time preference are more likely to drink alcohol and conditional upon drinking are more more likely to drink heavily. A policy maker with a better understanding of the determinants of time preference can design better policies that empower children to value their future well-being and thereby increase present healthy behavior.

Introduction

Empirical measurement of subjects' rates of time preference requires observations of subjects trading marginal future consumption against marginal present consumption. Although health behavior involves trades across time—pleasure increments now vs. health decrements later—drawing inference on rates of time preference observed in this context is complicated by a lack of common currency. The work of Thaler and Rosen (Thaler and Rosen 1976) and later Moore and Viscusi (Moore and Viscusi 1990; Viscusi and Moore 1989) established that observing labor markets where subjects trade higher wages in exchange for higher occupational fatality risk could be a useful means of measuring rates of time preference.

This paper extends the models of Moore and Viscusi by 1) Relaxing the assumption that workers are perfectly informed about occupational fatality risk; 2) Implementing a model of endogenous time preference. The extended model is applied to data from the National Longitudinal Survey of Youth (NLSY) to test Becker and Mulligan's (1997) suggestion that subjects adapt their rates of time preference as their future prospects and mental capacities develop. The extended model yields measures of time preference for each subject for each period. The validity of the time preference measures is explored by testing their relationship to alcohol use in the NLSY.

The term "time preference" refers to a subject's capacity to value future events [1]. The formation of consistent plans for intertemporal consumption places restrictions on the relationship

[1] The discount function $\beta(t)$ is defined as ratio $\beta(\tau) = U_t(X_{t+\tau})/U_t(X_t)$ where X_t is a good consumed at time t and $U_t(\ )$ is a utility function mapping consumption at various dates to utility at time t. The rate of time preference is defined as $-\log(\beta)$. Having no time preference is when the value of future satisfaction and the value of present satisfaction are equal in which case $\beta=1$ for all τ and time

between future satisfaction and current satisfaction. In theory, there is little reason for an optimal plan to change when nothing has changed but calendar time. Strotz described in 1956 the alternative conditions that were necessary for time consistency, 1) The discount function is an exponential function of the time which would elapse between the present and future consumption; or 2) The discount function is purely a function of calendar time; or 3) The discount function is a function of other values whose evolution is completely determined by calendar time(Strotz 1956).

The presumption of dynamic consistency is a staple of models of intertemporal choice. The typical assumption is constant exponential discounting [2]. The presumption is defended by suggesting that agents can have perfect foresight into the future dynamic evolution of their discount rates and employ precommitment strategies to enforce adherence to an optimal plan. In reality the question of whether human subjects have discount functions that support dynamically consistent planning is an empirical one. There has been little evidence to support constant exponential discounting (Ainslie and Haslam 1992; Cairns and van der Pol 1997; Loewenstein and Prelec 1992; Thaler 1981). This has led to greater consideration of alternative discounting models and a concern that dynamic inconsistency, if true, would invalidate basic assumptions

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preference = 0. Infinite time preference is when the subject does not value the future at all in which case $\beta=0$ and time preference $\rightarrow \infty$.

[2] A notable exception is Uzawa's demonstration that intertemporal models with discount functions that depend on the level of consumption require the assumption that β would fall with consumption (Uzawa 1968). To assume otherwise would lead to a degenerate case in which wealth accumulation approached infinity. Uzawa's assumption is counterintuitive to Blanchard and Fisher who suggest that one would expect that the wealthy have more future oriented time preference, $\beta \rightarrow 1$ (Blanchard and Fischer 1989).

required to apply rational choice models to environmental, health, and safety regulations (Loewenstein 1996).

One important form of dynamic inconsistency is the possibility that the cognitive limitations of teenagers might render them incapable of foreseeing changes in their time preference rates that will render them more concerned with their future welfare when they are older. If human development during the life course imposes dynamic inconsistency, it would offer an alternative explanation for the wave of regret that accompanies retrospection on the indiscretions of youth. Accounts of “regret” require an unforeseeable shift in the potential for harm from an activity. Orphanides and Zervos explain the regret from addictive behavior as rooted in unforeseeable shifts in the potential of young people to become addicted (Orphanides and Zervos 1994). Unforeseeable shifts in time preference would account for regret of non-addictive youthful behaviors such as unsafe sex, petty crime, and delinquency. Even if one granted that teenagers are fully informed about the risks from smoking, drinking, and unsafe sex, one would hesitate to press the welfare properties of a laissez faire approach to policies in this arena, if cognitive limitations made teenage risk choices dynamically inconsistent. A preliminary step to this line of inquiry is the estimation of valid measures of time preference during youth and across the lifecycle.

Methods

Measuring Time Preference in Individual Subjects in the NLSY.

Past attempts to obtain measures of time preference in individuals have used direct survey methods to assess self-reported time preference (Fuchs, 1982; Viscusi, Magat et al. 1987). But to study lifecycle variation in time preference one would require data on a sizable longitudinal cohort. Here, I identify individual measures of time preference in the context of panel data on individuals making repeated choices between money and danger[3]. Labor economists have devoted considerable attention to identifying the wage-risk tradeoff locus over the past four decades. See England (1992) for one review. This literature on compensating variations has recently been expanded to permit identification of the **discount rate** and the coefficient of risk aversion using readily available data on wages and occupational risk. (Moore and Viscusi 1988; Viscusi 1993; Viscusi and Moore 1989). It can be shown (Viscusi and Moore 1989) that in weighing the local danger premium that a subject must consider the lifetime value of the extra wages discounted by survival probability and discounted because of time preference.

The initial problem is the maximization of life time utility which can be written as:

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[3] Adam Smith is credited with the original observation that wages and danger are related. According to Smith, "In trades which are known to be very unwholesome, the wages of labour are always remarkably high" (Smith, 1776). Interestingly, the notion that risk attitude is different among adolescents also appears to originate with Smith in precisely the context of occupational choice. On pages 122-123 of Book 1 he writes, "The contempt of risk and the presumptuous hope of success, are in no period of life more active than at the age at which young people chose their professions…. A tender mother, among the inferior ranks of people, is often afraid to send her son to school at a sea-port town, lest the sight of the ships and the conversation and adventures of the sailors should entice him to go to sea."

$$[1] \quad \underset{p_j}{Max} V = U(w(p_j))[1 - p_j(p_j) - p_x(p_x)]\sum_{t=1}^{\infty}[b_j[1 - p_j(p_j) - p_x(p_x)]]^{t-1}$$

where:

U ( ) is the utility function
$p_j(\pi_j)$ is the perceived occupational fatality risk as a functional of <u>true</u> occupational risk π_j and $p_x(\pi_x)$ is the perceived lifetable fatality risk as a function of <u>true</u> lifetable risk, π_x.
$(1-p_j - p_x)^t$ is the perceived probability of surviving at least t periods given exposure to the perceived job fatality risk p_j, and lifetable driven perceived fatality risk p_x.
$w(\pi_j)$ is the payoff of the jth agent as a function of the occupation (and risk) selected
β_j is the discount rate of the jth agent

The model in Equation [1] modifies the Viscusi and Moore model by breaking the identification of perceived risk with actual risk. Imperfect knowledge of risk is modeled by incorporating the process whereby subjects perceive the fatality risks as "prospects" (Kahneman and Tversky 1979) "p" which have a functional association with the <u>true</u> risks now symbolized with Greek characters "π_j" and "π_x."

In this stripped down model, the subject selects once and for all their occupation which is parameterized only by its fatality risk. That choice determines the subject's wages according to $w(p_j)$ and the subject's survival $(1-p_j)$ for the subsequent t periods.

Noting that $\sum_{t=1}^{\infty}[b(1-p)]^{t-1}$ = 1/[1-β(1-p)], equation **Error! Not a valid link.**can be expressed as:

$$[2] \quad \underset{p_j}{Max} V = \frac{U(w(p_j))(1 - p_j - p_x)}{[1 - b(1 - p_j - p_x)]}$$

The first order condition is:

$$[3]\quad \frac{U'(w(p_j))(1-p_j-p_x)(\frac{\P\ w}{\P\ p_j})}{1-b\ (1\text{-}\mathrm{p}_\mathrm{j}\text{-}\mathrm{p}_\mathrm{x})} - \frac{U(w(p_j))\frac{\partial p_j}{\partial p_j}}{1-b\ (1\text{-}\mathrm{p}_\mathrm{j}\text{-}\mathrm{p}_\mathrm{x})} - \frac{b\ U(w(p_j))(1\text{-}\mathrm{p}_\mathrm{j}\text{-}\mathrm{p}_\mathrm{x})\frac{\partial p_j}{\partial p_j}}{[1-b\ (1\text{-}\mathrm{p}_\mathrm{j}\text{-}\mathrm{p}_\mathrm{x})]^2} = 0$$

Solving for $\frac{\P\ w}{\P\ p_j}$ yields

$$[4]\quad \frac{\P\ \mathrm{w}}{\P\ p_\mathrm{j}} = \frac{U(w(p_j))}{U'(w(p_j))}\left[\frac{1}{1-p} + \frac{b}{1-b\ (1\text{-}\mathrm{p})}\right]\left(\frac{\partial p_j}{\partial p_j}\right)$$

To estimate the model I will need to apply functional forms to U( ) and w (p_j) .

Viscusi and Moore (1989) explore the Constant Relative Risk Aversion (CRRA) utility functional form which can be parameterized as U=(w^c-1)/c. The limiting case of the CRRA utility function that turns out to be of interest for both exposition and estimation is c→0 in which case the CRRA utility function approaches the familiar logarithmic utility function. I thus assume:

[5] U = [log (w)]

With this functional form

$$[6]\quad \frac{U}{U'} = w\log(w)$$

and [4] can be rewritten as:

$$[7]\quad \frac{\P\ \mathrm{w}}{\P\ p_\mathrm{j}} = w\log(w)\left[\frac{1}{1-p_j-p_x} + \frac{b}{1-b\ (1\text{-}\mathrm{p}_\mathrm{j}\text{-}\mathrm{p}_\mathrm{x})}\right]\frac{\partial p_j}{\partial p_j}$$

Noting that dlogw/dπ = (1/w)*(dw/dπ) and rearranging yields:

$$[8]\left[\frac{1}{1-p_j-p_x} + \frac{b}{1-b\ (1\text{-}\mathrm{p}_\mathrm{j}\text{-}\mathrm{p}\ \ )}\right]^{-1}\frac{\P\ \log(\mathrm{w})}{\P\ p_\mathrm{j}} = \log(w)\frac{\partial p_j}{\partial p_j}$$

which can be simplified to yield

$$[9]\ \left[(1-p_j-p_x)-b\ \ (1-p_j-\mathrm{p}_x)^2\right]\frac{1}{\frac{\partial p_j}{\partial p_j}}\frac{\P\ \ l\mathrm{og(w)}}{\P\ \ p_{\mathrm{j}}}=\log(\mathrm{w})$$

Viscusi and Moore note that in actual practice the mortality risk is so low that the approximation 1-p=(1-p)^2 is warranted. With this approximation [9] becomes:

$$[10]\ \ \log(\mathrm{w})=(1-b)\left(\frac{(1-p_j-\mathrm{p}_x)}{\frac{\partial\,\mathrm{p_j}}{\partial\,p_{\mathrm{j}}}}\right)\frac{\P\ \log(\mathrm{w})}{\P\ \ p_{\mathrm{j}}}$$

Augmenting [10] with a vector of observed variables x_k and a vector of unobservable variables μ that may account for individual differences in tastes yields:

$$[11]\ \ \log(\mathrm{w})=(1-b)\left(\frac{(1-p_j-\mathrm{p}_x)}{\frac{\partial\,\mathrm{p_j}}{\partial\,p_{\mathrm{j}}}}\right)\frac{\P\ \log(\mathrm{w})}{\P\ \ p_{\mathrm{j}}}+\sum_k a_k x_k+m$$

With an approximation that the perception of job fatality is $p_j(\pi_j)$ is a linear function of the true job fatality, e.g. $p(\pi)= k*\pi$ then the relationship between wages and the market can be expressed as:

$$[12]\ \log w_j = C + (1-\beta_j)((1/k)-\pi_j-\pi_x)[dlogw_j/\ dp_j] + \alpha_j*X_j + \mu_{wj}$$

The intuition for Equation**Error! Not a valid link.** is straightforward. A positive coefficient on dlogw/dπ in Equation **Error! Not a valid link.** would indicate that higher subject wages are better explained by increases in "attractive local wage offers for dangerous jobs" when the subject faces a high probability of surviving his job risk and when the subject has a low discount rate (high time preference).

To interpret $(1-\beta_j)$ purely as a measure of time preference requires that statistical controls have been used for other mediators of the relationship between earnings and the danger premium. The foremost two confounders might be 1) economic desperation and 2) intrinsic risk attitude. I will control for desperation by using measures of family size, race, schooling, and gender as X_j variables. Parameterizing risk attitude within specialized functional forms like the constant relative risk aversion utility function offers one approach to making $(1-\beta_j)$ a cleaner measure of "time preference".

Note that if subjects perceive their risk perfectly k=1. With this assumption the coefficient of $(1-\pi_j-\pi_x)d\log w/d\pi_j$ is an unbiased measure of $(1-\beta_j)$. However what if subjects do not perceive risk perfectly? If subjects underperceive risks it will make 1/k large leading to data supporting a larger coefficient on the danger premium. The misspecification bias from assuming k=1 will lead to estimates of $1-\beta$ that are inordinately large. Because π_j and π_x are small, on the order of 10^{-5}, a good approximation for the magnitude of the misspecification bias is 1/k.

Eq **Error! Not a valid link.** imposes a restriction stating that other determinants of wages (X_j) such as schooling, age, gender etc. will independently shift the relationship between wages and wage-risk offers up and down (changes in intercept) but they will not alter its slope. I will loosen this restriction below. The italicized term in **Error! Not a valid link.** $d\log w_j / dp_j$ must be computed. To compute it, I will essentially use NLSY respondents as informants: They report to me their region, their wage, and their occupation. Using NIOSH's National Traumatic Occupational Fatality, data I convert reported occupation into a the fatality rate faced by that

NLSY respondent in exchange for the wage they were given. The statistical model implemented is a regression of the form:

$$[13] \quad \log w_j = C + \sum_k \left[\beta_{1k} (D_k{*}\pi_j) + \beta_{2k} (D_k{*}\pi_j^2) \right] + \beta_X X + \varepsilon$$

where D_k are dummy variables for region, and year.

Equation [**Error! Not a valid link.**is a representation of the hedonic market equilibrium for job safety (Kahn and Lang 1988). Controlling for other determinants of wages, the coefficient on occupational fatality will represent an equilibrium. Competition for workers between firms in a local labor market (determined by time and space) establishes the local balance between firms' willingness to invest in safer working conditions and their ability to simply raise wages to attract workers who are willing to jeopardize their survival for cash. Competition between workers for jobs establishes a similar balance between local workers' hedonic demand for better occupational survival poised against decrements in wages. As pointed out by Kahn and Lang (1988) were one to simply to regress log wage the X variables and on occupational risk *without the interactions*, the measures of $d\log w/d\pi$ would be endogenous in equation **Error! Not a valid link.**.

The X's are other determinants of wage. After running this regression I can use the results to compute as follows

$$[14] \quad \textit{danger premium} = \quad d\, \log w_j / dp_j = \sum_k (\beta_{1k})(D_k) + (\beta_{2k})(D_k)(\pi_j)$$

Because its computation depends only on regional and annual dummies introducing the predicted danger premium in **Error! Not a valid link.**will not introduce endogeneity bias.

Interacting $d\log w_j/dp_j$ with measures of $((1/k)-\pi_X-\pi_j)$ in which I set k=1 forms a new independent variable,

$Z_j = (1-\pi_X-\pi_j)^*$ *dlog* w_j/dp_j

To estimate $1-\beta_j$ I note that it seems plausible that in **Error! Not a valid link.** not just the intercept but the slope of the tradeoff between the danger premium and wages is a random coefficient that differs idiosyncratically in general patterns suggested by Becker and Mulligan's (1997) theory of endogenous time preference. Becker and Mulligan's theory offers an account of the determination of time preference by suggesting that individuals may alter their time preference in part by spending time and effort in forming mental pictures of future pleasures. They cite as examples of such time and effort activities like acquiring information through schooling, access to print media, and time spent with older persons, particularly parents. Subjects endogenously invest effort in these activities in proportion to the agreeableness of their future life prospects. For instance, Becker and Mulligan speculate that because church attendance offers repeated promise of future heavenly splendor it may lead attendants to endogenously develop ways to discount future splendor less.

I thus introduce a time preference determination equation inspired by Becker and Mulligan:

[15] $(1-\beta_j) = A + \gamma_a * \text{age} + \gamma_s * \text{schooling} + \gamma_g * \text{gender} + \gamma_r * \text{race} + \gamma_f * \text{father} + \gamma_c * \text{church} + \mu_{\beta j}$

which forms a system together with a shorthand version of equation [12]

[16] $\log w_{jt} = C + (1-\beta_{jt})^* Z_{jt} + \gamma_j^* X_{jt} + \mu_{wjt}$

<u>Data</u>

The National Longitudinal Survey of Youth (NLSY) offers panel data on a cohort of 12,686 young persons aged 14 to 21 in 1979 followed up until when they are aged 29 to 37 in 1994.

Table 2 describes the essential variables from NLSY to be included in the analysis of occupational choice. The National Traumatic Occupational Fatality data from NIOSH contains occupation specific fatality rates for US workers. These data have been merged with the NLSY data in order to perform the occupational risk study.

From Table 2 one can get a sense of the magnitude of selection bias. Attrition shrinks the sample size from 12,686 to 8,889, individuals who can be followed over 16 years. Attrition means that observations on the effect of higher age and later period are conditional on sample non-attrition.

Multilevel Model Estimation:

Equation system [**Error! Not a valid link.**and [16] is the multilevel system that will be estimated using the 15 waves of observations on however many of the 12,686 subjects in the NLSY report wages and occupations. Equation **Error! Not a valid link.** describes phenomena at the level of the individual across successive rounds, while equation [16] describes phenomena at the level of the individual within each round. Our approach is similar to that taken by Cairns and Van Der Pol (Cairns and van der Pol 1997) whose analysis of 2 rounds of survey data on intertemporal preferences using multilevel models decisively favored the multilevel approach over OLS.

Computing Individual Time Preference Measures

In addition to examining the coefficients for equation [15] to assess the effects of age, education, etc. on time preference, I will compute individual measures of β_j that incorporate estimates of the individual specific fixed effects $\mu_{\beta j}$. Estimates of β_j can be derived which maximize the joint likelihood function for the system set up by [15] and [16].

Obtaining Unbiased Standard Errors

The assumption that successive observations of the behavior of the same subject are truly independent is likely to be false. The error terms generated by the same subject in different years are likely to be correlated and will result in heteroskedasticity. The multilevel model takes into account precisely this form of heteroskedasticity.

Results

Determining the US Wage Risk Locus for 1979-1994

Table 3 presents estimates of equation [13] which is a first stage equation describing how the equilibrium wage risk locus shifts across local markets defined by quadrennial periods and 4 regions composed of roughly 15 states each. Although this is hardly the proper specification to examine the effects of the productivity controls, it builds confidence to review the ways in which the long established facts from labor economics are reaffirmed in Table 3.

- Marriage is associated with an increase in annualized wages of about 5% similar to previous findings (Daniel 1995)
- Each year of schooling is associated with an increase in annualized wages of about 5% similar to previous findings (Behrman et al. 1995; Schultz 1988)

- Workforce experience more than tenure in a given job is associated with increased wages similar to previous findings(Altonji 1987)
- Black race and female gender reduce annualized wages significantly see also(England et al. 1996)
- Union membership has a profound effect on wages (Leigh and Gill 1991)
- Occupational category controls appear to have reasonable effects as does job satisfaction.

All of the interaction variables are significant, and their coefficients are used together with equation [14] to obtain individualized estimates of the local danger premia. Job fatality interactions with the occupational dummies were not used because of the theoretical requirement that the dummies should mark off separate markets for occupational safety. By the same token, substantial migration *on the basis of danger premia*, would invalidate the use of the regional dummies to identify the occupational risk vs. wage market locus. There is no evidence to date for migration on the basis of geographical differences in compensating variations.

Determining the US Wage Determinants for 1979-1994

Table 4 presents parameters from the wage equation [16]. which is estimated jointly with the time preference determination equation [15] . The uninteracted covariates of wage have roughly the same coefficients in Table 4 as were noted in Table 3. The OLS estimates in Column I are derived by simply including the uninteracted survival adjusted danger premium, Z_j . The OLS estimates in Column II are derived by interacting each of the time preference determinants in Table 5 with survival adjusted local danger premium. OLS standard errors are adjusted for heteroskedasticity using White's correction. The Maximum Likelihood (multilevel) estimates in Column III are random coefficients estimates which append to the above interactions an interaction between survival adjusted danger premium and a person specific error term. This permits the model to account for unobservable idiosyncracies in a respondent's approach towards risky decisions.

A Tau test (Bryk and Raudenbush 1992) rejected the null of non-random coefficient on the danger premium, thus the multilevel model offers the preferred set of estimates as well as the preferred specification. The random coefficients produced by the maximum likelihood technique are displayed in Figure 1. Their distribution is approximately normal and centered around a mean of 5.3.

Can We Infer the Proper Discount Rate for Society?

As can be seen from equation [12] the interpretation of the danger premium coefficients as indicators of the discount rate require the assumption of perfect perception of occupational fatality rates (e.g. k=1). Although Viscusi and Moore (1989) are inclined to make this assumption, for both theoretical and empirical grounds, its validity is questionable. An empirical rejection of perfect risk perception in the NLSY can be based on the finding that coefficient on the danger premium at 5.3 is roughly 10 fold higher than it would need to be to place the discount rate in it's theoretical range of 0-1. If subjects underestimated the magnitude of their job fatality risk by at least a factor of 10 (1/k=0.1) , it would bring the mean estimate of β into the theoretical range. Unfortunately in the absence of information about the quality of risk perception, the empirical methods cannot deliver an absolute estimate of the population discount rate, nor of the value of future life. Those interested in obtaining estimates of these parameters for their own sake must either rely on unsupported assumptions or submit to disappointment. For our purposes explaining relative variation in discount rates is fully sufficient and we now turn to this task.

The Determination of Time Preference

Table 5 indicates that controlling for covariates young age, male sex, and white race are associated with an immediate time preference and or a less acute perception of risk. There is no evidence to support the conjecture of Becker and Mulligan that religious participation would encourage a less immediate time preference. We find no evidence to support the hypothesis that the presence of a father early in life fosters the capacity to consider one's future persistence and a less immediate time preference. Most surprisingly, we do not find evidence for an effect of schooling on time preference. According to the theory of endogenous preferences, schooling, by enhancing future earning potential would foster more attention to future great expectations. Part of the reason for our finding no effect of schooling may have to do with sample selection bias. The method used can only observe time preference in labor force participants. Labor force participants in their teens and early twenties have self selected out of college. Yet among older workers occupational choice estimates of college attenders are not comparable to non-attenders because of differences in work force experience. What more schooled subjects have in terms of time preference, they may lack in terms of risk perception and vice versa.

The Relationship of Time Preference to Risky Behavior

To assess the validity of time preference measures derived from studying compensating variations I test their performance in explaining differences in alcohol use. Measures of time preference have a prominent but infrequently tested association with health behavior in virtually all economic models of health behavior (Becker et al. 1990; Becker and Murphy 1988; Fuchs 1982; Grossman 1972; Grossman et al. 1994; Grossman et al. 1996; Ribar 1994). Thus, empirical measures of time preference can also be used to assess the strength of this association. Health-promoting behaviors that pay off in the future are more enjoyable to individuals who have lower time preference. The theory of rational addiction (Becker et al. 1991; Becker and Murphy 1988) highlights the way in which past consumption of an addictive good increases the marginal utility of current consumption. The theory predicts that individuals who greatly discount the future are more likely to become addicted and will be more sensitive to shifts in the price of the commodity. The policy implication of the theory is to use price shifts to alter addictive consumption of those who discount the future (assumed to be youth) and use public information campaigns to alter addictive consumption of those who value the future (assumed to be educated, adults) (Becker, 1991). Indeed purely on the basis of the assumption that youth discount the future more than do adults, the occurrence of greater price responsivity among youth has been credited as evidence in favor of the theory of rational addiction (Becker, et al. 1991; Grossman, et al. 1994). The evidence that youth are more price sensitive than adults is actually inconsistent (Chaloupka and Wechsler 1996).

To apply the relative time preference measures to explain relative differences in alcohol use rates, I first normalize the measures of β as Z-Scores then append them to a typical rational addiction equation such as $D_t=C+ \gamma_1 D_{t-1} + \gamma_2 D_{t+1} + \gamma_3 P_t + \varepsilon_t$ where D_t denotes drinking at time t and P denotes the price of a The two standard econometric problems to overcome in a rational addiction estimate are A. Skewness in the distribution of addiction in the population and B. Endogeneity of leads and lags of addictive consumption. To cope with skewness, I adopt a two part estimation strategy employing an initial probit for any alcohol use followed by estimates of heavy drinking among the sample of drinkers. To cope with endogeneity I adopt a two stage least squares approach using as instruments leads and lags of the real beer tax. The real beer tax data which is used as the price variable is available by state for 1979-1994 and is described in detail in {Chaloupka, 1995 #1017}.

Table 6 indicates that more patience (High β) is significantly related to both any alcohol use and among drinkers to the measure of the number of days of drunkenness in the past 30 days.

The three columns on the far right of Table 6 append to the rational addiction equation variable to control for other elements that are correlated with β by the very nature of its construction. Recall that β is computed according to Equation [15] as a linear combination of personal characteristics such as age, schooling, gender etc. plus an individual specific fixed effect that is computed by maximum likelihood. Appending the personal characteristics offers assurance that the effects found in Table 6 are not simply the effects of these potential confounders.

Discusssion

The finding of an aging effect on time preference leads to the possibility of dynamic inconsistency in planning one's consumption path. Finding that the measure of time preference is an independent predictor of both drinking and drunkenness suggests that lifecycle shifts in time preference are important for the study of health behavior. If adolescents fail to account for the likelihood that they will become more patient with age, they could make decisions that would have been rational had their rates of time preference remained constant, but which lead to regret in the face of life cycle changes in time preference.

Conclusion

I have expanded and refined a method pioneered by Viscusi and Moore (1989) to examine the response of subjects to various offers to trade their survival for money in the labor market. Because I demonstrate that the estimates obtained in this way depend critically upon the assumption of perfect risk perception, I do not share their conclusion that this method leads to unbiased estimates of the absolute rate of time preference in a population. I do believe that measuring and parameterizing these trades is a promising strategy for the study of relative differences in risky decision-making between subjects in a population. Although the measures obtained in this manner confound both time preference and the ability to accurately perceive fatality risks this confounding is neither conceptually nor empirically unappealing. Both factors are cognitive responses to contingent states of the world that dampen the full consideration of possibilities other than the present moment.

The investigation found that subjects become more patient with aging and that patience is independently correlated with less drinking and less drunkenness among drinkers. The presence of an aging effect on the discount rate leads to the possibility of dynamic inconsistency in planning intertemporal consumption paths. The fact that a “self” that starts down a consumption path could differ in an unforseeable way from one’s future “self” can lead to tragedy when the goods consumed have addictive or irreversibly harmful properties.

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Tables

Table 1

Table 1. Cohort structure of NLSY. Ages of subjects vs. Years of successive waves. The final row indicates the number of years in which the corresponding age group is observed.

Age	**14**	**15**	**16**	**17**	**18**	**19**	**20**	**21**	**22**	**23**	**24**	**25**	**26**	**27**	**28**	**29**	**30**	**31**	**32**	**34**	**35**	**36**	**37**
Year																							
79																							
80																							
81																							
82																							
83																							
84																							
85																							
86																							
87																							
88																							
99																							
90																							
91																							
92																							
93																							
94																							
obs	1	2	3	4	5	6	7	8	8	8	8	8	8	8	8	8	7	6	5	4	3	2	1

Table 2

Table 2. Variable Definitions and Means for Wage-Risk Estimates

Variable	**Years Available**	**Number of Obs in 1979**	**Number of Obs in 1994**	**Mean Conditional On No Attrition and Labor Force Participation**	**Source**
Endogenous					
Annual Real After Tax Wage	79-94	4657	7152	$9163	NLSY
Danger Premium (dlog w/dp_j) % Δ $ per unit Δ in Deaths/100K	79-94	4657	7152	0.0037	NTOF + NLSY
Exogenous Included					
Job Fatality Rate (p_j) (Deaths/100k)	90-91	4657	7152	4.6	NTOF
All Cause Mortality (Deaths/100k)	1990	12686	8889	134.6	NCHS
Years of Schooling	79-94	12686	8889	12.7	NLSY
Race (%African American)	79-94	12686	8889	23.4%	NLSY
Gender (%Female)	79-94	12686	8889	48.5%	NLSY
Age	79-94	12686	8889	25.5	NLSY
Number of Dependents (Incl Self)	79-94	12686	8889	3.24	NLSY
Union Status (% Members)	79-94	4852	6881	16%	NLSY
Exogenous (Instruments for Eq. [13])					
Period (4 quadrennial dummies)	79-94	12686	8889	NA	NLSY
Pd1 (1979-1982)	79-82	12686	NA	0.20	NLSY
Pd2 (1983-1986)	83-86	NA	NA	0.28	NLSY
Pd3 (1987-1990)	87-90	NA	NA	0.29	NLSY
Pd4 (1991-1994)	91-94	NA	8889	0.22	NLSY
Region (4 separate dummies)	79-94	12686	8889	NA	NLSY
Northeast	79-94	12686	8889	0.18	NLSY
South	79-94	12686	8889	0.38	NLSY
North Central	79-94	12686	8889	0.23	NLSY
West	79-94	12686	8889	0.19	NLSY

Table 3

Table 3. Identification of the Regional and Period Specific Danger Premia

Dependent Variable is Log of Annualized After Tax Earnings*

	Coefficient	t statistic
N	116,405	
R^2	0.4282	
Job Fatality * North East	0.0416	(16.24)
Job Fatality2 * North East	-0.0003	-(9.17)
Job Fatality * North Central	0.0289	(11.41)
Job Fatality2 * North Central	-0.0002	-(6.23)
Job Fatality * South	0.0306	(12.18)
Job Fatality2 * South	-0.0002	-(6.5)
Job Fatality * West	0.0416	(16.36)
Job Fatality2 * West	-0.0003	-(9.27)
Job Fatality * Period 1	-0.0311	-(12.24)
Job Fatality2 * Period 1	0.0002	(7.05)
Job Fatality * Period 2	-0.0316	-(12.54)
Job Fatality2 * Period 2	0.0002	(7.12)
Job Fatality * Period 3	-0.0275	-(10.87)
Job Fatality2 * Period 3	0.0002	(5.84)
Job Fatality * Period 4	-0.0254	-(10.)
Job Fatality2 * Period 4	0.0002	(5.36)
Highest Grade Completed	-0.0160	-(4.)
Highest Grade Completed^2	0.0025	(16.04)
Black	-0.0498	-(14.34)
Female	-0.1519	-(47.56)
Age	0.0725	(18.39)
Age^2	-0.0010	-(13.69)
Job Tenure	0.0010	(35.18)
Job Tenure^2	0.0000	-(19.86)
Experience in Job Market	0.0474	(25.41)
Experience^2	-0.0005	-(3.85)
Union Member	0.1725	(44.47)
Job Satisfaction	-0.0318	-(16.64)
Family Size	-0.0144	-(17.8)
Married	0.0490	(15.58)
Managerial**	0.2723	(51.58)
Technical Sales	0.1831	(43.88)
Operators & Fabricators	0.1258	(25.65)
Precision Production	0.2327	(40.68)
Farm/Forestry/Fishing	-0.0513	-(6.39)
Constant	7.5184	(139.17)

*Hourly Rate of Pay * 2000 Hours * (1-Tax Rate)
**Excluded Occupational Dummy: Service Industries
Period 1=79-82, Period 2=83-86, Period 3=87-90, Period 4=91-93

Table 4

Table 4. Estimating the Wage Equation in the NLSY

Dependent Variable is Log of Annualized After Tax Earnings*

	Column I		Column II		Column III	
	OLS		OLS		Unrestricted Maximum Likelihood	
	Coefficient	t-statistic	Coefficient	t statistic	Coefficient	t statistic
N	106,880		106,880		106,880	
Likelihood					-6.68E+04	
R^2	0.4383		0.4383			
Constant	7.5073	(87.82)	7.6120	(78.61)	7.7998	(130.7)
Survival Adjusted Local Danger Premium*	8.1250	(17.38)	*Distribution*		*Distribution Shown in Figure 2*	
Union Member	0.1731	(29.38)	0.1721	(29.44)	0.1631	(40.56)
Married	0.0539	(10.5)	0.0625	(10.56)	0.0536	(16.15)
Family Size	-0.0153	-(13.01)	-0.0164	-(12.6)	-0.0147	-(17.42)
Highest Grade Completed	-0.0226	-(2.48)	-0.0045	-(.32)	-0.0267	-(5.43)
Highest Grade^2	0.0027	(7.52)	0.0020	(3.58)	0.0029	(15.32)
Experience in Job Market	0.0460	(18.48)	0.0356	(12.97)	0.0404	(19.46)
Experience^2	-0.0003	-(1.98)	0.0000	(.01)	-0.0002	-(1.53)
Age	0.0787	(14.53)	0.0747	(11.17)	0.0702	(16.22)
Age^2	-0.0012	-(11.08)	-0.0013	-(9.62)	-0.0012	-(13.98)
Job Tenure	0.0010	(25.65)	0.0010	(26.83)	0.0009	(33.21)
Job Tenure^2	-9.36E-07	-(14.84)	-9.6E-07	-(15.15)	-1.00E-06	-(17.61)
Job Satisfaction	-0.0340	-(13.33)	-0.0339	-(13.37)	-0.0360	-(18.59)
Black	-0.0413	-(6.64)	-0.0306	-(4.46)	-0.0372	-(9.23)
Female	-0.1651	-(28.15)	-0.1731	-(25.1)	-0.1655	-(44.36)
North East			0.0223	(2.61)	-0.0098	-(1.24)
North Central			-0.0929	-(6.45)	-0.1445	-(17.26)
South			-0.0928	-(7.32)	-0.1373	-(18.46)
Managerial**	0.2601	(29.51)	0.2584	(29.41)	0.2317	(42.71)
Operators & Fabricators	0.1644	(23.01)	0.1556	(21.78)	0.1462	(29.7)
Technical Sales	0.1753	(25.91)	0.1784	(26.48)	0.1593	(36.77)
Precision Production	0.2711	(31.77)	0.2596	(30.31)	0.2322	(40.15)
Farm/Forestry/Fishing	0.0108	(.77)	-0.0222	-(1.57)	-0.0157	-(1.98)
Period 2			0.0383	(6.48)	0.0347	(6.76)
Period 3			0.1457	(12.71)	0.1545	(21.08)
Period 4			0.2179	(12.63)	0.2163	(22.61)

*Computed value of dlog(wage)/d job fatality from Table 3, interacted with survival probability from lifetable.

**Excluded Occupational Dummy: Service Industries

Period 1=79-82, Period 2=83-86, Period 3=87-90, Period 4=91-93

Table 5. Estimating Time Preference in The Level 2 Equation. Dependent Variable is Coefficient on Local Danger Premium in Table 4.

Dependent Variable is Coefficient on (1-π) dlogw/dπ

	OLS*		Maximum Likelihood	
	Coefficient	t statistic**	Coefficient	t statistic
Tau Test for Null of No Random Coefficients			618	
N	106880		106880	
Constant			-2.2082	-(1.81)
Lived with Father in 1979	0.5276	(.67)	1.2873	(1.53)
Religious Attendance	0.1763	(.77)	-0.3103	-(1.4)
Age	0.4323	(.47)	-6.3707	-(3.93)
Age^2	0.0032	(.2)	0.1365	(3.45)
Highest Grade Completed	-2.6616	-(1.45)	0.0435	(.03)
Highest $Grade^2$	0.1111	(1.67)	0.0482	(.86)
Black	-1.9782	-(2.04)	-3.6676	-(3.69)
Female	2.7100	(3.1)	1.2739	(1.65)
Period 2	2.3159	(2.55)	3.8277	(2.73)
Period 3	2.6211	(1.62)	7.7066	(2.09)
Period 4	0.2505	(.12)	-1.9350	-(.22)
Married	-1.3113	-(1.68)	0.9411	(.79)
Family Size	-0.0304	-(.16)	-0.4653	-(2.47)

* Model from Column II of Table 4

**OLS Standard errors adjusted for heteroskedasticity

Table 6. Rational Addiction Estimates of the Relationship Between Drinking and the Discount Rate

	Probit	OLS	2SLS	Probit	OLS	2SLS
Dependent Variable	**Any Drinking Last Month(1989)**	**Days Drank >6 Last Month 1989)**	**Days Drank >6 Last Month (1989)**	**Any Drinking (1989)**	**Days Drank >6 Last Month (1989)**	**Days Drank >6 Last Month (1989)**
Z Score of Discount Rate (β)	-0.043 (-2.062)**	-0.186 (-1.567)	-0.308 (-1.926)*	-0.041 (-1.876)*	-0.314 (-2.517)**	-0.462 (-2.765)***
Any Drinking Last Month (1984)	0.684 (14.688)***			0.606 (12.708)***		
Any Drinking Last Month (1994)	0.616 (13.950)			0.559 (12.416)***		
Days Drank >6 Last Month (1984) §		0.283 (16.323)***	-0.277 (-1.272)		0.261 (14.747)***	-0.283 (-1.274)
Days Drank >6 in Last Month (1994) §		0.355 (22.364)***	0.565 (2.618)***		0.343 (21.582)***	0.591 (2.928)***
Real Beer Tax (1989)	-0.121 (-3.065)***	-0.195 (-0.786)	-0.593 (-1.625)	-0.075 (-1.789)*	-0.218 (-0.856)	-0.528 (-1.512)
Age (Years)				-0.008 (-0.836)	-0.079 (-1.342)	-0.045 (-0.545)
Highest Grade (Years)				0.048 (4.885)***	-0.004 (-0.064)	0.0169 (0.229)
Married=1				-0.226 (-5.290)***	-1.026 (-4.135)***	-1.160 (-3.341)***
Female=1				-0.431 (-10.143)***	-1.338 (-5.168)***	-2.741 (-2.897)***
Black=1				-0.219 (-4.344)	0.1983 (0.628)	-0.380 (-0.747)
Lived with Father in 1979				-0.0408956 (-0.885)	-0.344 (-1.241)	-0.719 (-1.801)*
Religious Attendance				-0.005 (-0.414)	-0.120 (-1.577)	-0.278 (-2.071)**
Constant	-0.164 (-3.199)***	2.853 (12.380)***	5.827 (2.492)**	-0.043 (-0.141)	6.922 (3.792)***	9.753 (2.739)***
Adj or Pseudo R2	0.0917	0.2853	0.0313	0.1209	0.2968	0.0637
N	5118	2801	2277	5118	2801	2277

§ Denotes endogenous variables in 2SLS models. Instruments for 2SLS models are annual real beer taxes from 1979-1994.

Figure 1.

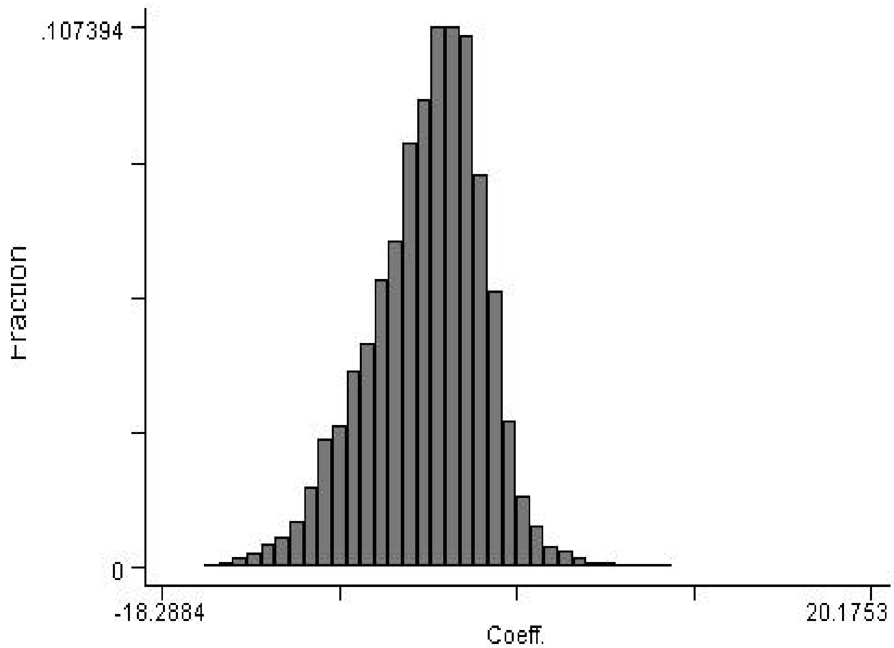

Figure 2
Maximum Likelihood Estimates